Media Coverage and IPO Pricing Around the World

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Abstract

We study how media coverage impacts pricing of IPOs around the world. Higher media coverage in the pre-IPO period leads to lower IPO initial returns. The effect is mitigated in countries with better financial reporting quality, greater shareholder rights protection, and more stringent media censorship, and for IPOs "certified" by reputable intermediaries, while it is amplified in countries with higher levels of media penetration and media trust. Further, IPOs with higher pre-IPO media coverage have lower ex-post price revision volatility. Our findings suggest that higher pre-IPO media coverage reduces information asymmetry among investors, leading to less underpriced IPOs.

JEL Classification: G10, G14, G15, G30

Keywords: Media coverage, IPO pricing, information asymmetry

I. Introduction

In recent years, the role of media in financial markets has become a focus of increasing attention by academics, practitioners, and regulators. While there is a general consensus that "media reporting can exert a large causal influence on financial markets" (Tetlock 2015, p. 703), the mechanisms through which that influence occurs as well as its implications for stock prices remain a subject of ongoing debate. One stream of research posits that, by disclosing and disseminating information to a broad population of investors, media coverage reduces information asymmetry and enhances informational efficiency of stock prices (Tetlock, Saar-Tsechansky, and Macskassy 2008; Drake, Guest, and Twedt 2014; Peress 2014; Twedt 2016). Another stream of research asserts that, by placing a firm in the spotlight of investor attention, media may exacerbate investor biases, causing stock prices to deviate from their fundamental values (Barber and Odean 2008; Dougal et al. 2012; Engelberg, Sasseville, and Williams 2012; Chen, Pantzalis, and Park 2013; Hillert, Jacobs, and Müller 2014).

Research examining the role of media in financial markets is predominantly U.S.-centric. In contrast, there has been little work published exploring the role of media in the global financial markets. As Griffin, Hirschey, and Kelly (2011, p. 3941) note: "Despite the perceived importance of the financial media, there has been little attempt to quantify its importance internationally or to understand why the impact of media varies across countries." Scarce evidence on the role of media in global financial markets constitutes an important gap in the literature given a sharp increase in the amount of capital raised in non-U.S. stock markets over the past two decades (Doidge, Karolyi, and Stulz 2013). We address this gap by examining the impact of media coverage on the pricing of initial public offerings (IPOs) around the world.

International IPOs provide an appealing platform for exploring the impact of media coverage on stock prices for several reasons. First, IPO firms are typically young, immature, and relatively informationally opaque (Ljungqvist 2007). Reflecting these features of IPO firms, both information asymmetry and investor limited attention—the mechanisms through which prior research suggests media influences stock price formation—play an important role in theories of IPO pricing. Second, legal institutions and the information environment—the "building blocks of efficiency" (Hung, Li, and Wang 2015)—both exhibit substantial variation across countries, thereby providing a powerful setting to explore the mechanisms through which media coverage affects stock prices. Third, to enhance interest in news stories, media coverage may be systematically biased toward stocks that experience substantial price changes (Shiller 2000; Bhattacharya et al. 2009), making causal inferences regarding the effect of media on stock prices problematic. Since IPOs do not have a share price history, the issue of reverse causality is mitigated in the IPO setting.

Prior research offers competing insights regarding the sign of the media coverage-IPO pricing relation. On the one hand, higher pre-IPO media coverage may reduce the "underpricing discount" in the offer price, resulting in lower IPO initial returns.[1] Prior research (Rock 1986; Benveniste and Spindt 1989; Benveniste and Wilhelm 1990; Spatt and Srivastava 1991; Michaely and Shaw 1994; Busaba and Chang 2010) shows that, in the presence of informed investors, IPO allocations have to involve underpriced stock. In the Rock (1986) model, some investors are assumed to be better informed about the true value of the shares on offer. This imposes a "winner's curse" on uninformed investors: they bid successfully for unattractive offerings, while in attractive offerings their demand is crowded out by informed investors. To prevent uninformed investors' withdrawal from the IPO market, new issues must be underpriced to allow uninformed investors to earn normal returns. Benveniste and Spindt (1989), Benveniste and Wilhelm (1990), and Spatt and Srivastava (1991) develop models in which underwriters elicit indications of interest from investors, which are then used in setting

[1] IPO initial return is measured as the percentage difference between the price at which the IPO shares are sold to investors (the offer price) and the first-day closing price of shares in the aftermarket.

the offer price. However, in the absence of inducements, informed investors have strong incentives to misrepresent positive information. To induce informed investors to reveal truthfully their information, the offer price needs to be discounted.

By disclosing and disseminating information, media leads to a more homogeneous distribution of information among investors, thereby reducing the information gap between informed and uninformed investors. Further, disclosure and dissemination of information reduces the level of ex-ante uncertainty regarding the value of the IPO firm. Since informed investors view information production as a call option on the IPO (Beatty and Ritter 1986), reduction of valuation uncertainty by the media mitigates investor incentives to engage in costly information gathering, thereby alleviating the impact of informational frictions among investors on IPO pricing. The above discussion suggests that high pre-IPO media coverage should reduce the magnitude of the underpricing in the offer price induced by informational frictions, and thus should be negatively associated with IPO initial returns.

An alternative perspective suggests that, by making an IPO firm more visible to retail investors in the aftermarket trading, high pre-IPO media coverage should lead to higher IPO initial returns. Attention is a scarce cognitive resource (Kahneman 1973). Consequently, there are cognitive and temporal limits to how much information an investor can process (Odean 1999; Barber and Odean 2008), making visibility of a firm's stock to investors an important attribute influencing their investment decisions. The impact of limited attention—and thus, the role of firm visibility—is particularly pronounced among individual (or retail) investors, who lack resources such as manpower and formal models to attend to and process information that is available to institutional investors (Battalio and Mendenhall 2005; Barber and Odean 2008). Further, limited attention has a stronger impact on buying, as individual investors search across thousands of stocks, than selling, where individual investors generally choose only from the set of stocks they own (Frieder and Subrahmanyam 2005; Barber and Odean 2008).

The role of media in priming investor attention could be particularly important in the context of IPOs, where most of the companies are relatively young and, therefore, less visible to investors. As discussed, by placing a firm in the spotlight of public discussion, media catches investors' attention (Engelberg and Parsons 2011; Solomon, Soltes, and Sosyura 2014; Hillert et al. 2014). Such an effect is particularly pronounced among the retail investors, who are net buyers of attention-grabbing stocks due to their lack of ability and resources to attend to and process information in a timely manner (Barber and Odean 2008). Consequently, by attracting retail investor attention to a new issue, higher pre-IPO media coverage may boost the price of a new issue in early aftermarket trading, resulting in higher IPO initial returns.[2,3]

In summary, insights from prior research offer competing theoretical predictions regarding the impact of media coverage on IPO pricing. The information asymmetry reduction mechanism suggests that higher pre-IPO media coverage alleviates informational frictions among the parties involved in an IPO, resulting in lower IPO initial returns. The visibility enhancement mechanism suggests that, by attracting retail investor attention to a new issue in the aftermarket trading, higher pre-IPO media coverage leads to higher IPO initial returns. Since it is not clear a priori which of the two mechanisms dominates, we frame the impact of media coverage on IPO pricing as an empirical question.

[2] The models of Derrien (2005) and Ljungqvist, Nanda, and Singh (2006) are particularly relevant to understanding the potential interplay between pre-IPO media coverage, IPO pricing, and investor attention. These models assume that there are two types of investor, informed institutional investors and individual investors who are assumed to be bullish (i.e., net buyers) at the time of the offering. The investment banker sets the offer price above its true value but below the valuation of individual investors. This allows the issuer to benefit from a higher valuation than appropriate, given the intrinsic value of the issue (as reflected in institutional investor valuations). In turn, institutional investors benefit from flipping their shares to individual investors in the early aftermarket trading. Hence, by promoting a new issue to retail investors, high pre-IPO media coverage may increase IPO valuation in the aftermarket relative to the final offer price, resulting in higher IPO initial returns. Said differently, the theoretical mechanisms advanced by Derrien (2005) and Ljungqvist et al. (2006) stipulate that the impact of investor limited attention—and, by inference, the role of pre-IPO media coverage in attracting investor attention—will be reflected in the higher spread between the final offer price (set in the primary market) and the price at which the new issue is traded in the early aftermarket trading.

[3] Relatedly, Da, Engelberg, and Gao (2011) measure the effect of investor attention on IPO underpricing through Google searches of firms during the IPO filing period. They find that high levels of Google searches are associated with higher initial returns.

Using a sample of 11,716 IPOs across 39 countries for the period 2000–2014, we find that higher pre-IPO media coverage is associated with lower IPO initial returns. This finding is in line with the information asymmetry reduction mechanism. The documented effect of media coverage is economically meaningful: in our sample, a one-standard-deviation increase in pre-IPO media coverage, on average, is associated with a decrease of 508 basis points in IPO initial returns. The documented effect of media coverage is robust to alternative model and sample specifications and holds in country-by-country and year-by-year regression analyses.

The IPO initial return and the level of pre-IPO media coverage could both be driven by firm or underwriter attributes not accounted for in our analysis. Hence, endogeneity is a potential concern in our setting. To address this concern, we conduct two tests. First, we adopt a quasi-natural experiment approach, using national media strikes as exogenous shocks to media coverage (Peress 2014). Second, we estimate our model using an instrumental variable estimation approach. Following Dai, Parwada, and Zhang (2015), we use a firm's geographic proximity to a Dow Jones branch as an instrument for media coverage. The negative association between media coverage and IPO initial returns holds in both analyses. Therefore, we conclude that our findings are unlikely to be driven by endogenous effects.

We further explore cross-sectional patterns in the strength of the documented media coverage–IPO pricing relation. If our arguments are valid, the documented effect of media coverage on IPO pricing should be mitigated (amplified) in settings where the role of media in generating and disseminating information among investors is less (more) salient.

We commence our cross-sectional analysis by examining the role of country-specific financial reporting quality and shareholder rights protection in the media coverage–IPO pricing relation. Higher quality of financial reporting system enhances the transparency of financial reports (Bhattacharya et al. 2003) and thus should reduce information asymmetry associated with an IPO. The legal rules of jurisdictions in which securities are issued and the quality of their enforcement are important determinants of what rights securities holders have and how well these rights are protected, which in turn determines their willingness to finance firms (LaPorta et al. 1998). Therefore, we reason that investor exposure to information asymmetry-related risks—and thus, the role of media in reducing information asymmetry—should be mitigated in countries with better financial reporting quality and stronger legal protection of investors. Consistent with this, we find that negative association between media coverage and IPO initial returns is mitigated in countries with greater accounting conservatism, higher levels of anti-director shareholder rights, and stronger securities laws, while it is amplified in countries with higher earnings opacity and civil law countries.[4]

Next, we examine the impact of media penetration, censorship, and trust on the relation between media coverage and IPO pricing. Media penetration facilitates dissemination of news through media channels among investors (Zingales 2000; Dyck and Zingales 2004). Further, media is more effective in disseminating news when it has greater credibility with the public (Dyck, Volchkova, and Zingales 2008; You, Zhang, and Zhang 2017). In contrast, media censorship obstructs the access to valuable information and stifles independent criticism and analysis (Geddess and Zaller 1989; Shadmehr and Bernhardt 2015). In our setting, these insights suggest that the role of media coverage in reducing information asymmetry—and thus, the documented effect of media coverage—should be stronger in countries with higher level of media penetration and media trust and weaker in countries with higher level of media censorship. Our findings lend support to these predictions.

We further explore the impact of IPO certification on the media coverage–IPO pricing relation. Prior research suggests that the presence of prestigious intermediaries (e.g., prestigious underwriters and/or reputable auditors) reduces information asymmetry faced by

[4] Civil law countries generally have weaker legal investor protection compared to common law countries (LaPorta et al. 1998). Therefore, investor exposure to information asymmetry-related risks is expected to be more pronounced in these countries.

investors (Booth and Smith 1986; Titman and Trueman 1986; Carter and Manaster 1990; Michaely and Shaw 1994; Amihud, Hauser and Kirsh 2003). In a similar vein, the presence of venture capitalists as investors in a firm going public plays an important certification role for a new issue (Megginson and Weiss 1991; Loughran and Ritter 2004). Therefore, the role of media coverage in reducing informational frictions should be mitigated for IPOs with stronger certification characteristics. In line with these arguments, we find that the documented effect of media coverage is mitigated for IPOs underwritten by investment banks with strong reputation, IPOs audited by Big 4 auditing firms, and IPOs backed by venture capitalist investors.

To further gauge the mechanism driving our findings, we supplement our analysis with three sets of tests. First, we explore the interplay between information asymmetry on the primary IPO market, media coverage, and IPO underpricing. If the information asymmetry reduction channel plays a role in our setting, we should observe (i) a negative association between media coverage and the level of information asymmetry on the primary IPO market, and (ii) a positive association between the level of information asymmetry on the primary IPO market and IPO initial returns. Using ex-post price revision volatility as a proxy for information asymmetry on the primary IPO market, we find support for both predictions.

We also examine whether the effect of media coverage on IPO initial returns varies depending on news content, type, tone, and the timing of the article. We reason that if the documented effect of media coverage occurs through the information asymmetry reduction mechanism, such an effect should be particularly pronounced for articles that focus on an IPO firm's earnings, as firm earnings play a pivotal role in investor valuation of an IPO (Brau and Fawcett 2006; Willenborg, Wu, and Yang 2015). We also reason that the documented effect of media coverage should be stronger for the full article category (i.e., articles that provide news analysis) compared to other, less informative, types of articles (e.g., news flashes and short press releases). Further, we expect the role of media in reducing information asymmetry—and thus, the documented effect of media coverage—to be weaker when there is greater variance in the tone of the media (i.e., greater disagreement in media regarding IPO firm prospects). Finally, we expect the documented effect of media coverage to be stronger for news articles published closer to the listing date—i.e., when investor interest in the IPO is more likely to "heat-up" (Liu et al. 2014). Our results lend support to these predictions.

We conclude our analysis by reconciling our findings with results from U.S.-based studies (Cook, Kieschnick, and Van Ness 2006; Liu, Sherman, and Zhang 2017). Focusing predominantly on the pre-2000 period, these studies document a positive association between media coverage and IPO initial returns on the U.S. market. When considered within the context of these studies, our findings suggest that the media coverage-IPO initial returns relationship on the U.S. market has undergone a structural shift in the post-2000 period. To examine the empirical validity of this conjecture, we supplement our cross-country analysis by comparing the relationship between media coverage and IPO initial returns on the U.S. market in the pre-2000 period with the one in the post-2000 period (the sample period used in our study). Consistent with Cook et al. (2006) and Liu et al. (2017), we document a positive association between media coverage and IPO initial returns in the pre-2000 period. However, we also find that the sign of this relationship becomes significantly negative in the post-2000 period, consistent with the information asymmetry reduction channel. The results of additional analyses indicate that the shift in the sign is not driven by a decline in the information technology IPOs following the collapse of the dot-com bubble and is robust to using alternative measures of media coverage. Collectively, these results suggest that the visibility enhancement channel has dominated the media coverage-IPO pricing relationship on the U.S. market in the pre-2000 period, whereas information asymmetry reduction channel dominates this relationship in the post-2000 period.[5]

We contribute to the literature along several important dimensions. Our first contribution is to the literature examining the impact of media on capital markets (see Tetlock 2015 for a review). So far, this strand of literature has been predominantly U.S.-centric, thereby offering limited insight regarding the role of media in global financial markets (Griffin et al. 2011). We advance this literature by generating comprehensive evidence on the role of media in reducing information asymmetry within the context of international IPOs. In this context, our study relates to Griffin et al. (2011), who also investigate the role of media in global financial markets. While Griffin et al. (2011) focus on the interplay between media and stock market volatility around news events, we examine the impact of media on IPO pricing.

Second, we contribute to the literature on IPO pricing (see Ljungqvist 2007 for review). Understanding the determinants of IPO pricing in global markets is particularly important given a substantial increase in the share of world IPO activity by non-U.S. firms (Doidge et al. 2013). Our findings suggest that higher media coverage leads to lower IPO initial returns, consistent with the notion that media coverage reduces information asymmetry associated with a new issue. Our findings also suggest that media coverage-IPO pricing relation on the U.S. market has undergone a structural shift, and future studies may delve into mechanisms driving this shift. This suggests an avenue for future research.

Third, our study contributes to the growing literature on the role of financial reporting quality and legal institutions in capital markets (LaPorta et al. 1998; Djankov et al. 2008; Spamann 2010; He and Hu 2014). Our findings suggest that the role of media in reducing information asymmetry among investors is particularly pronounced in countries with poorer quality financial reporting and weak shareholder rights protections. In this context, our study

[5] We discuss potential reasons for this structural shift in Section VII.

addresses the Ljungqvist (2007) call to examine the determinants of IPO pricing in a cross-country setting. These results also cast media as an important informal institution that alleviates informational frictions in settings where formal institutions offer limited protection to investors, thus potentially carrying policy implications for regulators.

II. Sample and Variables

A. Sample

We obtain media coverage data from RavenPack, a leading global media database widely used in recent finance and accounting research (e.g., Shroff, Verdi, and Yu 2014; Dai et al. 2015; Dang, Moshirian, and Zhang 2015; Twedt 2016; Bushman, Williams, and Wittenberg-Moerman 2017; You et al. 2017). Starting in 2000, RavenPack has gathered news articles around the world from three major sources: (1) Dow Jones newswires, regional editions of *The Wall Street Journal*, and *Barron's*; (2) business publishers, national and local news, blog sites, and government and regulatory updates; and (3) press releases and regulatory, corporate, and news services, including PR Newswire, the CNW Group (formerly the Canadian News Wire), and the Regulatory News Service.

We obtain IPO data from the Security Data Company's (SDC) Platinum New Issue Database. Firm-level financial information and stock returns data are obtained from Datastream and Worldscope. Data on country-level economic development and quality of listing stock exchange are obtained from the World Bank's World Development Indicator database. Following prior literature (Cook et al. 2006; Liu, Sherman, and Zhang 2014), we exclude unit offers (IPOs with warrants), closed-end funds, real estate investment trusts, and limited partnerships. We follow prior research on international IPOs (e.g., Lin, Pukthuanthong, and Walker 2013) by excluding issues with a converted offer price below US$1.00. Further, we require IPO firms to have information in Datastream or Worldscope at least in the IPO year. We also require each country in our sample to have at least 10 IPOs. Our final sample consists of 11,716 IPOs from 39 countries spanning the period 2000–2014. To mitigate the effect of potential outliers, we winsorize all variables (except for dummy variables) at both the upper and lower 1-percentile.

B. Variables

Our dependent variable is IPO first-day return (*First-day return*). Following prior studies (e.g., Ellul and Pagano 2006; Colak et al. 2017), we calculate *First-day return* as the first-day closing price of an IPO minus its offer price scaled by the offer price. Our explanatory variable of interest is media coverage in the pre-IPO period (*Media coverage*). In constructing this variable, we follow Liu et al.(2014) and count the total number of news articles about the IPO firm within the 30-day period prior to the IPO date.[6] We use the log-transformed number of news articles as our measure of pre-IPO media coverage, defined as the natural logarithm of 1 plus the number of news articles about the IPO firm in a 30-day window prior to the IPO date. RavenPack assigns a relevance score for each news article (ranging from 0 to 100), indicating how strong the news article is related to a specific firm. Following prior studies (e.g., Drake et al. 2014; Dang et al. 2015), we focus on the news articles with a relevance score of 100 to ensure that these articles are primarily about the firm under discussion. We further utilize the event similarity key and RavenPack story identification code to identify the original news releases and exclude duplicate entries. If an IPO firm has no reported news articles during the 30-day period, we set the number of news articles to zero following Cook et al. (2006).

Our selection of control variables follows prior literature (e.g., Ellul and Pagano 2006; Colak et al. 2017). *Firm size* is calculated as the natural logarithm of total assets of the IPO firm. *Profitability* is defined as earnings before interest and taxes divided by total assets. *Leverage* is measured as the ratio of total debt over total assets. *Market-to-book* is calculated

[6] We focus on 30-days prior to the IPO date because investor interest in the IPO is more likely to "heat-up" in a short window leading up to the firm's listing (Liu et al. 2014).

as market value of assets divided by the book value of assets. *Asset turnover* is calculated as sales divided by total assets. *Bookbuilding* is a dummy variable equal to 1 if the IPO is conducted using a bookbuilding method, and zero otherwise. Following Ellul and Pagano (2006) and Doidge et al. (2013), we also control for the state of the economy and the level of capital market development in the country where an IPO takes place. We include *GDP per capita growth*, measured as growth in annual GDP per capita, *Market size*, measured as the ratio of annual total value of stocks traded to GDP, and *Market turnover*, measured as the aggregate stock market turnover ratio. Detailed variable definitions are provided in the Appendix.

C. Summary Statistics

Table 1 reports the sample distribution and summary statistics of the variables in the analysis. Panel A presents the distribution of IPOs across the 39 countries in our sample. The panel shows that China has the largest number of IPOs, followed by U.S., Japan, U.K., and Australia. The panel also shows that China has the highest average first-day return (70.4%), while Portugal has the lowest average first-day return (-8.9%). In terms of pre-IPO media coverage, the panel shows that Taiwan has the largest average number of news articles (24.553), while Ireland has the smallest average number of news articles (6.409).

[Insert Table 1 about here]

Panel B presents the distribution of IPOs across time in our sample. The panel shows that the global IPO market reached its peak in 2007 in terms of the number of IPOs and declined gradually thereafter. This observation is in line with recent industry reports (e.g., Ernst and Young 2016). Average IPO first-day return was highest in 2010 (72.0%) and lowest in 2000 (20.4%). The average number of articles covering an IPO firm in the 30-day window prior to the listing day reached its peak in 2009 (21.858) and was lowest in 2000 (10.24).

Panel C presents the summary statistics of the variables. The panel shows that the average *First-day return* in our sample is 38.7%.[7] The average number of articles is 14.058 and the average value of the log-transformed number of articles (i.e., average *Media coverage*) is 1.755. The average IPO firm in our sample has *Firm size* of 4.811, *Profitability* of 0.042, and *Leverage* of 0.121. The highest Variance Inflation Factor among the explanatory variables is 1.26, suggesting that multicollinearity is not a concern in our analysis.

III. Empirical Results

A. Univariate Analysis

We begin with a univariate analysis of the media coverage-IPO initial returns relation. We divide sample IPO firms into 11 groups based on the number of news articles about an IPO firm appearing within the 30-day window prior to the IPO date. We assign IPOs with no news articles to Group 0 and then divide the remaining IPOs into 10 equal groups based on the number of news articles. Group 1 has the lowest number of news articles while Group 10 has the highest. We then calculate the average IPO first-day return for each of the 11 groups and plot the results in Figure 1. Figure 1 shows that the average IPO first-day return in Group 0 is 59.9%, whereas the average IPO first-day return in Group 10 is only 17.6%. The difference between the average returns in the two groups is statistically significant (p-value<0.01). The figure also shows that average IPO first-day return decreases monotonically when moving from Group 0 to Group 10. As discussed earlier, this pattern is consistent with media coverage reducing information asymmetry among investors, which, in turn, results in a lower degree of IPO underpricing. Next, we explore whether this effect persists in a multivariate framework.

[Insert Figure 1 about here]

B. Baseline Regression Analysis

[7] Average IPO first-day return in our sample is higher than that reported in prior studies (e.g., Lin et al. 2013; Boulton, Smart, and Zutter 2010). This is because these studies do not include Chinese IPOs, which have exceptionally high IPO first-day return (Tian 2011).

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Here, we examine the relation between pre-IPO media coverage and IPO first-day return using regression analysis. The specification of our baseline regression is as follows:

$$\begin{aligned}
&First\ day\ return_{i,j,t} = \alpha + \beta_1 Media\ coverage_{i,j,t} + \beta_2 Firm\ size_{i,j,t} \\
&+\beta_3 Profitability_{i,j,t} + \beta_4 Leverage_{i,j,t} + \beta_5 Market - to - book_{i,j,t} \\
&+\beta_6 Asset\ turnover_{i,j,t} + \beta_7 Book\ building_{i,j,t} + \beta_8 GDP\ per\ capita_{j,t} \\
&+\beta_9 Market\ size_{j,t} + \beta_{10} Market\ turnover_{j,t} + \sum FE + \varepsilon_{i,j,t}
\end{aligned} \tag{1}$$

where *i* denotes IPO firm, *j* denotes country, *t* denotes year, $\sum FE$ denotes fixed effects, and ε is the error term. The model is estimated with year, country, and industry fixed effects included. We use the 10 industry classifications as detailed on Kenneth French's website (http://mba.tuck.dartmouth.edu/pages/faculty/ken.french/). The model is estimated using pooled ordinary least squares (OLS) with standard errors adjusted for heteroskedasticity.[8]

We report the results in Table 2 using a set of nested models. Column (1) presents the results with industry, year, and country fixed effects but without control variables; column (2) presents the results with firm-level control variables; column (3) presents the results of our full baseline model, which also includes country-level control variables. In all three specifications, the coefficient of *Media coverage* is significantly negative (highest *p*-value<0.01), suggesting that higher media coverage is associated with lower IPO first-day return. These results confirm our findings from the univariate analysis and are consistent with media coverage reducing the magnitude of IPO underpricing. The coefficient of *Media coverage* in Column (3) suggests that a one-standard-deviation increase in *Media coverage* (1.373) reduces IPO first-day return by 0.037×1.373 = 0.0508 or 5.08 percentage points which constitutes a 13.14% reduction compared to the mean. Hence, we conclude that the effect of media coverage on IPO pricing is economically meaningful. The results for the control variables are largely consistent with prior studies (e.g., Ellul and Pagano 2006; Doidge et al. 2013; Colak et al.2017): IPO first-day return

[8] In (untabulated) robustness test, we repeat all our analyses with standard errors clustered at the country level and obtain qualitatively similar results.

is positively related to *Firm size*, *Profitability*, *Leverage*, *Market-to-book*, and *Market size*, while negatively related to *Asset turnover* and *Bookbuilding*.

[Insert Table 2 about here]

C. Country- and Year-Specific Regressions

Since our sample includes IPOs from multiple countries and multiple years, a potential concern is that the documented effect of media coverage is driven by IPOs from a particular country or a particular year. To examine this issue, we conduct two tests.

First, we estimate our baseline regression model for each country. Each country-specific regression includes all the control variables from Equation (1), industry fixed effects, and year fixed effects. To account for the appropriate level of degrees of freedom, we perform tests only for countries with at least 50 IPOs in our sample, which results in 24 country-by-country regressions. We present the results in Panel A of Table 3. For brevity, we report only the coefficient of *Media coverage*. The table shows that the coefficient of *Media coverage* is negative in 20 out of the 24 countries. Further, despite a sharp reduction in sample size available to estimate country-specific regression models, the coefficient of *Media coverage* is significantly negative at the 10% level or better in 17 countries. These results suggest that the documented effect of media coverage is not limited to IPOs from a particular country.

Second, we estimate our baseline regression model for each year. Each year-specific regression includes all the IPO-firm-level control variables from Equation (1), industry fixed effects, and country fixed effects. The results are reported in Panel B of Table 3 showing that the coefficient of *Media coverage* is significantly negative at the 10% level or better in 12 out of 15 year-specific regressions.[9] Based on these results, we conclude that the documented effect of media coverage is not driven by IPO observations from a particular year.

[9] The coefficient of *Media coverage* is not significant in 2000, 2002, and 2008, possibly due to high turbulence in stock markets during the dot-com bubble and the Global Financial Crisis periods.

[Insert Table 3 about here]

IV. Robustness Tests

To further assess the robustness of our findings, we conduct a battery of sensitivity tests. For brevity, we report only the coefficient of *Media coverage*. Control variables and fixed effects are included in all regressions but are not tabulated.

First, we explore the robustness of our findings to alternative time windows over which IPO initial returns are measured. Ljungqvist (2007) notes that in less developed capital markets, or in the presence of daily volatility limits, aftermarket prices may take some time before they equilibrate supply and demand. In our setting, these insights suggest that using first-day IPO return may underestimate the impact of media coverage. To examine this issue, we re-estimate our baseline model twice: once using IPO returns measured over one week following listing day and once using IPO returns measured over two weeks following listing day (Ellul and Pagano 2006; Lin et al. 2013). The results are reported in Panel A of Table 4 showing that the coefficient of *Media coverage* is significantly negative in both tests and that the magnitude of the coefficient is similar to that of our baseline model. Hence, we conclude that our findings are not sensitive to the choice of time window over which IPO initial returns are measured.Second, we examine the sensitivity of our results to alternative measures of pre-IPO media coverage. As outlined earlier, we construct our media coverage measure by counting the number of news articles over the 30-day window prior to the IPO date. To ensure that our results are not driven by specific choice of pre-IPO time window, we re-estimate our baseline model twice: once with media coverage constructed using a 60-day pre-IPO window, and once with media coverage constructed using a 90-day pre-IPO window. Last, we construct the media coverage measure using data from Factiva to verify that our findings are not driven by the choice of the media database. The results, reported in Panel B of Table 4, show that the coefficient of *Media coverage* remains significantly negative in all tests.

Third, we augment Equation (1) with additional control variables to mitigate concern that our findings are driven by some variable (or variables) not included in the baseline model. Specifically, we include *Advertising intensity*, calculated as the ratio of advertising expenditure to sales (Chemmanur and Yan 2017), *IPO float*, measured as the percentage of regular shares issued by the firm to the public and available to trade (Brennan and Franks 1997), and *Hot issue market*, measured as average IPO initial return for IPOs issued over the three months prior to a firm's IPO (Bradley and Jordan 2002). We further include *IPO size*, measured as the ratio of total IPO proceeds to the book value of total assets, *IPO age*, measured as the log-transformed number of years from the year when the firm was founded up to the year of listing, and *Cash balance*, measured as the ratio of cash holdings of the IPO firm scaled by total assets (Ellul and Pagano 2006; Lin et al. 2013). The results are reported in Panel C of Table 4 showing that the coefficient of *Media coverage* remains significantly negative in this estimation.[10]

Fourth, we examine the robustness of our results to exclusion of IPOs with no information in RavenPack during the 30-day window. As discussed, we follow prior research by setting the number of news articles to zero if an IPO has no information in RavenPack. To ensure that our findings are not driven by this practice, we re-estimate our baseline regression model after excluding these IPOs from the sample. The results, reported in Panel D of Table 4, show that the coefficient of *Media coverage* remains significantly negative in this estimation.

[Insert Table 4 about here]

We also consider the possibility of confounding effects in the media coverage-IPO pricing relationship, where the information asymmetry channel in the primary market dominates the attention channel in the secondary market. To examine this possibility, we decompose IPO initial return into its primary market component (i.e., the difference between

[10] We do not include these additional control variables in our main design because of their effects on our sample size. Our baseline sample has 11,716 observations (see Table 2). When additional controls are included, the sample size reduces to 7,481 observations.

first-day opening price and the final offer price scaled by the final offer price) and the secondary market component (i.e., the difference between first-day closing price and first-day opening price scaled by first-day opening price). We then re-estimate our baseline model twice, using each of the two return components as the dependent variable, respectively. Due to data availability, the sample size for this test is reduced to 2,232 observations. The results of this test are reported in Panel E of Table 4. The coefficient of *Media coverage* is significantly negative in the primary market return regression, consistent with the information asymmetry reduction channel, while insignificant in the secondary market return regression, thus, providing no evidence on the attention channel.

Last, we control for potential non-monotonicity and non-linearity in the media coverage-IPO pricing relation (e.g., Pollock and Rindova 2003).[11] To that end, we estimate our baseline model using semiparametric technique (Robinson 1988). This technique estimates the link function between the dependent variable (in our setting, *First-day return*) and the explanatory variable of interest (in our setting, *Media coverage*) nonparametrically, thereby imposing no assumptions (such as monotonicity or linearity) on the functional form of the examined relation. The results of semiparametric estimation are reported in Figure 2. The figure depicts monotonic and, generally, linear negative relation between media coverage and IPO initial returns, with the upper tail of *Media coverage* variable being the only exception, where the slope of the curve seems to flatten.[12] Consistent with this, the (untabulated) results indicate that allowing for non-parametric (i.e., assumption-free) specification of the media coverage-

[11] The examined relation may be non-monotonic if there is a "tipping point" in the level of media coverage where the visibility enhancement mechanism overtakes the information asymmetry reduction mechanism, causing the sign of the relation to change from negative to positive. The non-linearity in the media coverage-IPO initial returns relation may arise if, after media coverage reaches a certain level, the role of additional media coverage in reducing information asymmetry—and thus, it's effect on IPO initial returns—weakens.

[12] Importantly, the flexibility of semiparametric estimation comes at cost of lower estimation accuracy as reflected in larger standard errors (e.g., Pagan and Ullah 1999). An increase in standard errors is particularly pronounced at the tails of the distribution of the explanatory variable, where the observations are scarce. Therefore, the non-linearity at the upper tail of *Media coverage* reported in Figure 2 may also be driven by sampling issues associated with low estimation accuracy at the tails of the *Media coverage* variable.

IPO initial returns relation leads to no discernible improvement in the model performance compared to a more parsimonious linear regression specification. To further control for potential non-linearity, we examine the robustness of our findings to (i) excluding the observations in the upper tail of *Media coverage,* and (ii) using decile-ranking of *Media coverage*. The (untabulated) results remain qualitatively similar to those reported in the paper.

[Insert Figure 2 about here]

V. Endogeneity

A potential concern is that our findings could be driven by some attribute (or attributes) correlated with both the level of pre-IPO media coverage and IPO first-day return. While the results reported earlier partially alleviate this concern, they do not rule it out completely. To further examine this issue, we conduct two tests.

In the first test, we use national media strikes as exogenous shocks to media coverage (Peress 2014). Consistent with Peress (2014), we focus on strikes that affect the press on a national scale and involve the media sector only. These strikes are called by journalists, print, or distribution workers, and typically relate to their profession's economic conditions (i.e., employment, pay, pensions, tax breaks, state subsidies, and other benefits). Therefore, these media strikes are not driven by—and thus, are exogenous to—stock market movements and/or economy-wide conditions during the period of the strike (Peress 2014).

We use the list of media strikes reported by Peress (2014, Table 1). During our sample period, we identify 31 eligible national media strikes that meet the criteria discussed in the previous paragraph. To perform the test, we first identify 581 IPOs that were conducted in the countries where the media strikes took place during our sample period. Then, we assign each of these IPOs to either the "strike" group or the "non-strike" group. The "strike" group includes the IPOs with at least one media strike taking place in the same country during the 30-day window prior to the IPO date. This results in 62 IPOs in this group. The "non-strike" group includes all the remaining 519 IPOs. Finally, we construct a strike dummy (*Strike*) equal to 1 if the IPO is in the "strike" group, and zero if it is in the "non-strike" group.

To assess the validity of using media strikes as exogenous shocks to media coverage in our setting, we first test whether there is a significant difference in media coverage between the "strike" and "non-strike" groups. The results, presented in Panel A of Table 5, show that the mean number of news articles for IPOs in the "strike" ("non-strike") group is 9.726 (14.202), and the difference between the two groups is statistically significant (p-value<0.04). This finding is consistent with media strikes reducing the level of pre-IPO media coverage. Next, we examine the impact of media strikes on IPO first-day return. To perform the test, we estimate the same regression specification as in Equation (1), but with the strike dummy as the explanatory variable of interest. The results, reported in Panel B of Table 5, show that the coefficient of the strike dummy is significantly positive (p-value<0.03), suggesting that a lower level of media coverage resulting from media strikes leads to higher IPO first-day return. These findings support the causal effect of media coverage on IPO first-day return.[13]

[Insert Table 5 about here]

Second, we estimate our baseline regression model using an instrumental variable approach. The media coverage of a firm is dependent on the distance between the firm and news outlets (Gurun and Butler 2012). Therefore, we follow Dai et al. (2015) by using a firm's geographic proximity to a Dow Jones branch as an instrumental variable. Specifically, we construct a categorical variable *Proximity to DJ branch*, which equals 2 if the IPO firm is headquartered in a metropolitan area with at least one Dow Jones news branch, 1 if the IPO firm is headquartered in a country with at least one Dow Jones news branch but in a metropolitan area without any Dow Jones news branch, and zero otherwise. We obtain

[13] To ensure that our results do not reflect economic effects of media strikes on IPO firms in media-related industries, we repeat the analysis in Table 5 after excluding these IPOs from our sample. In our sample, we have seven IPOs in media-related industries (SIC codes 2711, 3663, and 4833). The (untabulated) results indicate that exclusion of these IPOs from our sample has no material impact on our findings.

information about the location of Dow Jones news branches around the world from the Dow Jones website (https://www.dowjones.com/).

In our setting, a valid instrumental variable should meet the following two selection criteria (Larcker and Rusticus 2010): (1) it should be significantly correlated with *Media coverage,* and (2) it should not be correlated with the residuals of our baseline regression model (i.e., it should not have a direct effect on IPO first-day return). Consistent with the first criterion, the results reported in Column (1) of Table 6 show that the coefficient of *Proximity to DJ branch* in the *Media coverage* regression is significantly positive. The (untabulated) partial *F*-statistic for the exclusion test of *Proximity to DJ branch* from the *Media coverage* regression is above the critical value of 8.96 (Stock, Wright, and Yogo 2002), suggesting that the weak instrument issue is not a concern in our setting. Consistent with the second criterion, there is no ex-ante reason to suggest that, after controlling for *Media coverage*, the distance between the IPO firm headquarters and a Dow Jones news branch has a direct effect on IPO first-day return. Therefore, we conclude that *Proximity to DJ branch* is a valid instrument in our setting. The results of instrumental variable estimation are reported in Column (2) of Table 6 showing that the coefficient of *Media coverage* is significantly negative (p-value<0.01). This result provides further support that our findings are not driven by endogenous effects.

[Insert Table 6 about here]

VI. Cross-Sectional Tests

A. Moderating Effects of Country-Level Institutions

In this section, we examine the effects of country-level financial reporting quality, shareholder rights protection, and emerging market status on the media coverage–IPO pricing relation. Higher quality financial reporting alleviates information asymmetry among investors and mitigates associated agency problems (Mahoney 1995). In terms of shareholder rights protection, legal rules and the quality of their enforcement at a country-level where an IPO is carried out determine what rights securities holders have and how well these rights are executed, which in turn determines the willingness of investors to finance firms (LaPorta et al. 1998). Given the above discussion, we reason that the role of media in alleviating information asymmetry-related risks—and thus, the documented effect of media coverage—is mitigated in countries with higher quality financial disclosure and stronger legal protection of investors.

We begin by examining the effect of country-specific financial reporting quality on the media coverage–IPO pricing relation. Following Boulton et al. (2011, 2017), we use *Earnings opacity* and *Accounting conservatism* scores of the country where the IPO takes place as our measures of country-specific financial reporting quality.[14] A higher value of *Earnings opacity* indicates lower quality of financial disclosure, whilst a higher value of *Accounting conservatism* indicates higher quality of financial disclosure. To examine the effect of country-level financial reporting quality on the media coverage–IPO pricing relation, we modify our baseline model to include the interaction term between the two measures and *Media coverage*, respectively. The results are reported in columns (1) and (2) of Table 7 showing that the coefficient of *Media coverage*×*Earnings opacity* is significantly negative and the coefficient of *Media coverage*×*Accounting conservatism* is significantly positive (highest p-value<0.01). These results support our prediction that the effect of media coverage on IPO pricing is mitigated in countries with higher financial reporting quality.

Next, we examine the effect of shareholder rights protection on the media coverage–IPO pricing relation. We employ three measures of shareholder rights protection. Our first measure is *Security law*, which measures the average number of country-specific disclosure

[14] Earnings opacity score for each country captures the extent of earnings aggressiveness, loss avoidance, and earnings smoothing. Higher earnings opacity score reflects lower quality of information conveyed by firms' earnings to investors (Boulton et al. 2011). Accounting conservatism refers to accounting practices, policies, and tendencies through which firms reported net asset values are understated relative to their market values. Boulton et al. (2017) show that IPOs are underpriced less in markets where accounting conservatism is more prevalent, consistent with accounting conservatism providing investors better information about IPO firms. The earnings opacity score is not available for Chile, Poland, and Russia. The accounting conservatism score is not available for China, Poland, and Russia. Therefore, we exclude IPOs from these countries in tests using the two measures.

requirements by stock exchange, liability standards, and public enforcement of legal contracts (LaPorta et al. 2006). We obtain the data required for the construction of *Security law* from the World Bank's Doing Business Indicators. Our second measure is *Shareholder right*s, which is the anti-director self-dealing rights index of the country where the IPO occurs. We obtain the index values from Djankov et al. (2008) and Spamann (2010). Higher values for *Security law* and *Shareholder right*s indicate better investor protection. Our third measure is *Civil law*, which is a dummy variable equal to one if the IPO occurs in a civil law country in our sample, and zero otherwise.[15] In civil law countries, the interests of minority shareholders are not well protected, thereby exposing them to greater risk of managerial expropriation (LaPorta et al. 1998). We interact each of these variables with *Media coverage* and include the interaction terms separately in Equation (1). The results are reported in columns (3)-(5) of Table 7 showing that the coefficients of *Media coverage*×*Security law* and *Media coverage*×*Shareholder rights* are both significantly positive, while the coefficient of *Media coverage*×*Civil law* is significantly negative (highest *p*-value<0.01). These results support our prediction that the effect of media coverage on IPO pricing is mitigated in countries with stronger levels of shareholder rights protection.

Last, we examine the effect of emerging market status on the media coverage–IPO pricing relation. To that end, we modify our baseline model to include the interaction term between the emerging market dummy (*Emerging*) and *Media coverage*.[16] The results are reported in column (6) of Table 7 showing that the coefficient of the *Media coverage*×*Emerging* variable is significantly negative (*p*-value<0.01), suggesting that the effect of media coverage on IPO pricing is stronger in the emerging markets. This result is

[15] We identify civil law countries following LaPorta et al. (1998). In our sample, the civil law countries are Argentina, Austria, Belgium, Brazil, Chile, China, Denmark, Finland, France, Germany, Greece, Indonesia, Italy, Japan, Mexico, Netherland, Norway, Philippines, Poland, Portugal, Russia, Spain, Sweden, Switzerland, South Korea, Taiwan, and Turkey.

[16] Emerging markets in our sample include Argentina, Brazil, Chile, China, India, Indonesia, Mexico, Malaysia, Philippines, Poland, Russia, Thailand, Turkey and South Africa.

consistent with emerging markets, typically, having lower quality institutions and weaker investor protection than developed markets (Chen, Hope, Li, and Wang 2011), which elevates the role of media in mitigating information asymmetry-related risks. [17]

[Insert Table 7 about here]

B. Moderating Effects of Country-Level Media Characteristics

In this section, we explore the impact of country-level media penetration, media censorship, and media trust on the relation between media coverage and IPO pricing. Prior research (e.g., Zingales 2000; Dyck and Zingales 2004; Mullainathan and Shleifer 2005) suggests that the breadth of news dissemination through media depends on the level of media penetration among the investors—i.e., the extent to which investors have access to and use the media channels to obtain the information. These insights suggest that the role of media in disseminating information among investors—and thus, the documented effect of media coverage—should be stronger in countries with greater level of media penetration.

We employ two measures of media penetration. Our first measure, *Newspaper users*, captures country-level penetration of press (Dyck and Zingales 2004) and is calculated as the proportion of people subscribing to newspapers in a country. Our second measure, *Internet users*, captures country-level internet penetration (Boulton et al. 2015), and is calculated as the proportion of people subscribing to the internet in a country. We obtain the data for the construction of these variables from World Bank's World Development Indicators database. To test our prediction, we interact each of the two variables with *Media coverage* and include the interaction terms separately in the regression specification in Equation (1). The results are reported in columns (1) and (2) of Table 8 showing that the coefficients of *Media*

[17] We do not control for country fixed effects in regressions reported in Tables 7 and 8 to avoid perfect collinearity between the stand-alone effect of the country-specific time-invariant moderator and country fixed effect dummies. As an (untabulated) robustness test, we re-estimate the models in Tables 7 and 8 with country fixed effects included and without stand-alone effect of the country-specific moderator. Using this alternative specification has no material impact on our findings.

coverage×*Newspaper users* and *Media coverage*×*Internet users* are both significantly negative (highest *p*-value=0.031). These results support our prediction that the effect of media coverage on IPO pricing is stronger in countries with greater level of media penetration.

Next, we examine the impact of country-level media censorship on the media coverage–IPO pricing relation. By controlling the flow of news and information to the public, media censorship obstructs the access to valuable information and stifles independent criticism and analysis (Geddess and Zaller 1989; Shadmehr and Bernhardt 2015). Based on these insights, we expect media censorship to hinder the role of media as an informational intermediary in the IPO process. Therefore, we forward that the documented effect of media coverage on IPO pricing should be weaker in countries with more stringent media censorship.

We employ two measures of media censorship. Our first measure, *Press censorship*, captures country-level censorship of the press and reflects the degree of print, broadcast, and digital media freedom in a country. Our second measure, *Internet censorship*, captures country-level censorship of the internet and reflects the strength of legal and ownership control rights over internet service providers, limits on the on-line content, and frequency of user rights violations. We obtain the data on media censorship from Freedom House (https://freedomhouse.org/). For each of the two variables, larger value indicates more stringent censorship. The (untabulated) results indicate that both variables are significantly correlated with country-level measures of financial disclosure quality and investor protection in our sample. To isolate the effect of censorship, we orthogonalize our censorship measures with respect to these variables. We then interact each of the two censorship variables with *Media coverage* and include the interaction terms separately in in Equation (1). The results are reported in columns (3) and (4) of Table 8 showing that the coefficients of *Media coverage*×*Press censorship* and *Media coverage*×*Internet censorship* are both significantly positive (highest p-value<0.01). These results support our prediction that the effect of media coverage on IPO pricing is weaker in countries with more stringent media censorship.

Last, we investigate the effect of country-level media trust on the media coverage–IPO pricing relation. Prior studies (e.g., Dyck et al. 2008; You et al. 2017) document that media is more effective in disseminating news when it has greater trust of the public. Therefore, we expect the role of media in reducing information asymmetry—and thus, the documented effect of media coverage—to be amplified in countries with greater level of media trust.

We employ two measures of media trust. Our first measure, *Confidence in Press*, captures country-level confidence in the information disseminated by the press. Our second measure, *Confidence in TV News,* captures country-level confidence in the information broadcasted on TV. We obtain data of the two media trust measures from the World Value Survey (http://www.worldvaluessurvey.org/wvs.jsp). For each of the two measures, higher value indicates greater level of public trust in media. Analogous to our censorship measures, we orthogonalize media trust variables with respect to country-level measures of financial disclosure quality and investor protection. We then interact each of the two media trust variables with *Media coverage* and include the interaction terms separately in the regression specification in Equation (1). The results are reported in columns (5) and (6) of Table 8 showing that the coefficients of *Media coverage×Confidence in Press* and *Media coverage×Confidence in TV News* are both significantly negative (highest p-value<0.01). These results support our prediction that media trust amplifies the effect of media coverage on IPO pricing.

[Insert Table 8 about here]

C. Moderating Effect of IPO Certification

In this section, we examine the effect of IPO certification on the relation between media coverage and IPO pricing. The IPO process involves a substantial degree of information asymmetry between insiders and outside investors. To mitigate information asymmetry, outside investors attempt to obtain information about the IPO firm from various sources. In particular, investors are thought to infer the quality of the IPO firm based on whether the IPO firm is backed by venture capital firms (Megginson and Weiss 1991; Loughran and Ritter 2004), whether the firm is audited by a high-quality auditor prior to the IPO (Menon and Williams 1991), or whether there is a high-quality underwriter that underwrites the IPO (Carter and Manaster 1990). Said differently, the presence of venture capital firms and/or reputable intermediaries "certifies" the quality of the issue by mitigating information asymmetry and the associated agency concerns. Therefore, we expect the role of media in alleviating informational frictions—and thus, the documented effect of media coverage—to be mitigated for IPOs with greater levels of certification by the third parties associated with the IPO.

We employ three measures of IPO certification (Carter and Manaster 1990; Megginson and Weiss 1991; Menon and Williams 1991; Loughran and Ritter 2004). Our first measure is the venture capital indicator variable (*VC back*), which is equal to 1 if the IPO firm is backed by a venture capital firm, and zero otherwise. Our second measure is the Big 4 auditor indicator variable (*Big 4 auditor*), which is equal to 1 if the IPO firm is audited by one of the Big 4 auditing firms, and zero otherwise. Our third measure is the reputable underwriter indicator variable (*Underwriter*), which is equal to 1 if the investment bank underwriting the IPO is in the top quartile based on combined IPO proceeds, and zero otherwise. We interact each of the three variables with *Media coverage* and include the interaction terms separately in Equation (1). The results are reported in Table 9 showing that the coefficient of *Media coverage*×*VC back* is significantly positive, and so are the coefficients of *Media coverage*×*Big 4 auditor* and *Media coverage*×*Underwriter* (highest *p*-value=0.012). These results are consistent with strong certification characteristics mitigating the effect of media coverage on IPO pricing.

[Insert Table 9 about here]

VII. Additional Analysis

A. Media Coverage and Information Asymmetry on the IPO Primary Market

As discussed, we argue that higher media coverage mitigates information asymmetry among IPO investors, thereby reducing the degree of IPO underpricing aimed to compensate investors for the information asymmetry-related risks. In this section, we test the empirical relevance of the proposed mechanism in our setting.

Since information asymmetry is unobservable, providing direct evidence on the hypothesized mechanism requires a use of a proxy. Information asymmetry induces heterogeneity in investors' beliefs, resulting in higher volatility of stock prices (Shalen 1993; Grundy and Kim 2002). Building on these studies, we employ ex-post price revision volatility as a proxy for the level of information asymmetry among the IPO investors. Drawing on prior literature (Andersen and Bollerslev 1998; Granger and Sin 2000; Forsberg and Ghysels 2007), we construct two measures of ex-post price revision volatility, *Revision_Vol_SqRes* and *Revision_Vol_AbsRes*. We calculate *Revision_Vol_SqRes* (*Revision_Vol_AbsRes*) as the square (absolute value) of the price revision regression residual. The price revision regression residuals are obtained from regressing price revisions on media coverage and the control variables, capturing the deviations of price revisions from their expected values. If the hypothesized information asymmetry channel is valid, we should observe (i) a positive association between ex-post price revision volatility and IPO first-day returns, and (ii) a negative association between media coverage and ex-post price revision volatility.

First, we examine the association between ex-post price revision volatility and IPO first-day returns. To that end, we regress *First-day return* on each of the two measures ex-post price revision volatility and the control variables. To ensure that the examined relationship is separate from the partial adjustment effect documented in prior IPO literature (e.g., Hanley 1993; Bradley and Jordan 2002), we also control for the run-ups in IPO prices in this analysis (untabulated for brevity). The results are reported in columns (1) and (2) of Table 10 showing that the two measures load positively in the regressions (highest p-value<0.01). These results provide support for our first conjecture—i.e., that, in our sample, higher level of information asymmetry is associated with larger underpricing discount. Next, we examine the association between media coverage and ex post price revision volatility. To that end, we regress each of the two measure of ex-post price revision volatility on *Media coverage* and the control variables. The results are reported in columns (3) and (4) of Table 10 showing that *Media coverage* loads negatively in both regressions (highest p-value<0.01). These results provide support for our second conjecture—i.e., that, in our sample, higher level of media coverage is associated with lower level of information asymmetry.

[Insert Table 10 about here]

B. Effects of the Content, Type, Tone, and Timing of the Article

In this section, we examine whether the impact of media coverage on IPO initial returns varies depending on news content, type, tone, and timing of the article. Prior research suggests that a firm's earnings play a pivotal role in investor assessment and valuation of an IPO (Brau and Fawcett 2006; Willenborg et al. 2015). Therefore, if the documented effect of media coverage occurs through the information asymmetry reduction mechanism, such an effect should be stronger for news articles that focus on IPO firm earnings. To test this prediction, we use the RavenPack classification scheme to divide the news articles into three categories: (i) *Earnings news,* which includes news articles about the IPO firm's earnings (e.g., earnings releases and earnings revisions), (ii) *IPO news,* which includes articles with news specific to the IPO (e.g., news about IPO approval, delay, and extension), and (iii) *Other news*, which includes news articles that do not belong to either of the two aforementioned categories. We count the number of news articles in each of these three categories and include the three (log-transformed) counts in Equation (1). The results, reported in column (1) of Table 11, show that the coefficient of media coverage is significantly negative for each of the three categories.

Moreover, the coefficient of *Earnings news* is significantly larger (in absolute terms) than the coefficients of *IPO news* and *Other news* (highest *p*-value=0.02), consistent with the IPO firm earnings news playing a dominant role in reducing information asymmetry among investors.

We further explore whether the documented effect of media coverage varies depending on article type. RavenPack classifies news articles into five types: full article, hot news flash, news flash, press release, and tabular material. In contrast to the other four types, news articles in the full article category have both a headline and textual material in the body. Since full articles are more informative, we expect the effect of media coverage on IPO pricing to be stronger for news articles in the full article category compared to news articles in other categories. To test this prediction, we divide the total number of news articles covering the firm into two categories: (i) *Full articles*, which includes news articles classified by RavenPack as full articles, and (ii) *Other articles*, which includes news articles in the other four categories. We count the number of news articles in each of these two categories and include the (log-transformed) counts in Equation (1). The results, reported in column (2) of Table 11, show that the coefficient of media coverage is significantly negative for both categories. Moreover, the coefficient of *Full article* is significantly larger (in absolute terms) than the coefficient of *Other articles* (*p*-value=0.01). These results are consistent with the more informative news articles playing a more pronounced role in mitigating information asymmetry among IPO investors.

We also explore whether the tone of media has an impact on the documented effect of media coverage. We reason that the role of media in reducing information asymmetry should be weaker when there is greater variance in the tone of the media (i.e., greater disagreement in media regarding IPO firm prospects). We employ the RavenPack event sentiment score (ESS) to capture the tone of the articles. The score ranges from 0 (extremely negative tone) to 100 (extremely positive tone) where 50 represents neutral tone. Following Bushman et al. (2017), we assign a value of 1 if the ESS is greater than 50, a value of -1 if the ESS is less than 50, and a value of 0 if the ESS is equal to 50. We calculate *Media variation* as the standard deviation of the transformed ESS of all the news articles about the IPO firm within the 30-day window prior to the IPO date and include the interaction term between *Media coverage* and *Media variation* in Equation (1). The results are reported in column (3) of Table 11 showing that the coefficient of *Media coverage*×*Media variation* is significantly positive (p-value<0.01). This result is consistent with greater level of divergence in media tone weakening the role of media in reducing information asymmetry among investors.

We further examine whether sign of the media coverage-IPO initial returns relation is contingent upon the direction of the media tone. Information asymmetry reduction channel suggests that news articles convey information to investors, thereby reducing information asymmetry associated with the IPO. Hence, information asymmetry reduction channel predicts a negative association between media coverage and IPO underpricing irrespective of the tone of media. In contrast, investor attraction channel suggests that articles with positive (negative) tone induce positive (negative) sentiment among investors. This perspective suggests that the sign of the media coverage-IPO initial returns relation should be positive (negative) for the articles with positive (negative) tone. To examine which of these two channels operates in our setting, we classify news articles into *Positive news* (articles with ESS above 50), *Negative news* (articles with ESS below 50), and *Neutral news* (articles with ESS equal to 50) categories. We count the number of news articles in each of these categories and include the three (log-transformed) counts in Equation (1). The results, reported in column (4) of Table 11, show that the coefficient of media coverage is significantly negative for all three categories (highest p-value=0.064), supporting the information asymmetry reduction channel. The coefficients of *Positive news* and *Negative news* are both significantly larger (in absolute terms) than that of *Neutral news* (highest p-value=0.01), suggesting that the tone of media impacts the extent to which media coverage reduces IPO underpricing.

Last, we investigate whether the documented effect of media coverage varies with the timing of news articles. We divide the 30-day pre-IPO window into three 10-day time windows: [-10,-1], [-20,-11], and [-30,-21]. Then, we count the number of news articles in each time window and include the (log-transformed) counts in Equation (1). The results, reported in column (5) of Table 11, show that the coefficient of media coverage is significantly negative for the [-10,-1] and [-20,-11] time windows (highest *p*-value<0.01) and is negative and marginally significant for the [-30,-21] time window (*p*-value=0.089). Further, the coefficient of media coverage monotonically increases (in absolute terms) as we move from the [-30,-21] to [-10,-1] time window (*p*-value of the difference test=0.044). These results are consistent with the notion that news articles published closer to the listing date (i.e., when investor interest in the IPO is more likely to "heat-up" (Liu et al. 2014) play a more pronounced role in mitigating information asymmetry among IPO investors and thus, have a stronger effect on IPO underpricing.

[Insert Table 11 about here]

C. Reconciliation with Prior U.S. Evidence

We conclude by reconciling our findings with results from U.S.-based studies (Cook et al. 2006; Liu et al. 2017). Focusing predominantly on the pre-2000 period and using alternative databases (such as Factiva or LexisNexis) to capture media coverage, these studies document a positive association between media coverage and IPO initial returns on the U.S. market, consistent with the visibility enhancement channel.[18] When considered within the context of these studies, our findings suggest that either (i) the sign of the media coverage-IPO returns relationship on the U.S. market is contingent on the choice of database used to construct a

[18] As discussed in Section II, we follow extensive research in finance and accounting by using RavenPack database to construct our measure of media coverage. In our setting, RavenPack offers at least two key advantages over other media databases (such as Factiva or LexixNexis): (i) it provides a much more comprehensive coverage of non-U.S. IPOs and (ii) it classifies articles based on the topic and the tone of the article, which we use to facilitate identification of theoretical mechanism driving our findings (see discussion in Section VII.B). Since RavenPack begins its coverage in 2000, our analysis covers the post-2000 period.

media coverage measure, or (ii) the media coverage-IPO returns relationship on the U.S. market has undergone a structural shift in the post-2000 period. To examine the empirical relevance of these scenarios, in this section we focus our analysis on the U.S. market.

To facilitate comparison with prior studies, we employ Factiva in searching news articles about U.S. IPO firms prior to the IPO date (Liu et al. 2014; 2017). Our U.S. sample consists of 2,348 IPOs that have information available in Factiva spanning the period 1993-2014. We partition this sample to estimate two regressions: one using the 1993-2000 sample period (the sample period used by Cook et al. 2006, hereafter "pre-2000 period") and one using the 2001-2014 period (hereafter "post-2000 period"). The results are reported in Panel A of Table 12. Consistent with Cook et al. (2006) and Liu et al. (2017), we document a positive association between media coverage and IPO initial returns in the pre-2000 period (*p*-value<0.01). However, we also find that the sign of this relationship becomes significantly negative in the post-2000 period (*p*-value<0.01). This result is consistent with our findings in Panel A of Table 3 which use the RavenPack-based media coverage measure for U.S. IPOs. The results of robustness tests, reported in Panel B of Table 12, show that the documented shift in the sign is not driven by a decline in the information technology IPOs following the collapse of the dot-com bubble and is robust to using alternative measures of media coverage.

[Insert Table 12 about here]

Collectively, the results reported in Table 12 suggest that the visibility enhancement channel has shaped the media coverage-IPO pricing relationship on the U.S. market in the pre-2000 period, whereas information asymmetry reduction channel dominates this relationship in the post-2000 period. This structural shift potentially reflects the confounding effect of two factors. First, there has been a significant decline in the U.S. IPO market activity since 2000 (Gao, Ritter, and Zhu 2013). Prior research suggests that the role of firm visibility is particularly salient in a setting where investors have to interpret multiple signals coming from a large number of firms (Hirshleifer, Lim, and Teoh 2009; Frederickson and Zolotoy 2016). By inference, the role of media in attracting investors' attention to a new issue is expected to be more salient in highly active IPO market, with multiple firms going public within a short period of time. Therefore, a decline in the IPO activity may have weakened the visibility enhancement role of media in the U.S. IPO market.[19] Second, major financial turmoils in the post-2000 period—such as the collapse of the dot-com bubble, prominent accounting scandals (e.g., Enron and Worldcom), and the Great Recession—have eroded investors' trust in corporations and capital markets (Agrawal and Chadha 2005; Lins et al. 2017). A decline of public trust in formal corporate governance and regulatory mechanisms may have elevated the role of media as an informal ("extra-legal") monitoring institution (Klausner 2005; Borden 2007) in reducing information-related risks through disclosure and dissemination of information to investors.

VIII. Conclusions

We examine the effect of media coverage on IPO pricing in markets around the world. We find that higher pre-IPO media coverage, on average, leads to lower IPO initial returns. The documented effect of media coverage is mitigated in countries with higher financial reporting quality, stronger shareholder rights protection, and more stringent media censorship, while amplified in countries with higher levels of media penetration and media trust. Further, the effect of media coverage is mitigated for IPOs with stronger certification characteristics and IPOs with greater variance in the tone of the media. We also find that IPOs with higher pre-IPO media coverage have lower ex-post price revision volatility. Taken together, our findings suggest that higher pre-IPO media coverage reduces information asymmetry among investors, leading to less underpriced IPOs.

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[19] Gao et al. (2013) report that during 1980–2000, an average of 310 companies per year went public in the U.S. Since 2000, the average has been only 99 IPOs per year. They further report that the drop was especially precipitous among small firms—i.e., firms that are less visible to investors (Frederickson and Zolotoy 2016).

We contribute to the emerging literature that examines the impact of media on capital markets (see Tetlock 2015 for a review). So far, this literature has been predominantly U.S.-centric, thus offering limited insights on the role of media in the global markets (Griffin et al. 2011). We also contribute to the literature on IPO pricing (see Ljungqvist 2007 for a review) by showing that higher level of pre-IPO media coverage results in less underpriced IPOs. We also contribute to the literature that examines the role of financial disclosure standards and legal institutions in capital markets (LaPorta et al. 1998; Djankov et al. 2008; Spamann 2010; He and Hu 2014). Collectively, our findings cast media as an important "informational intermediary" in IPO pricing in markets around the world, and thus should be of interest to academic researchers, investors, and regulators.

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Appendix: Variable Definitions

Variable	Definition
Accounting conservatism	Country-specific accounting conservatism score based on Boulton et al. (2017).
Advertising intensity	Advertising expenses divided by sales of the IPO firm at the time of listing.
Age	Log transformation of 1 plus the difference in years since the firm was established up to the year of listing.
Asset turnover	Sales divided by total assets of the IPO firm at the time of listing.
Big 4 auditor	Dummy variable equal to 1 if the IPO firm is audited by a Big 4 auditing firm, and 0 otherwise.
Bookbuilding	Dummy variable equal to 1 if IPO uses bookbuilding method, and 0 otherwise.
Cash balance	Cash and short-term investments divided by total assets of the IPO firm at time of listing.
Civil law	Dummy variable equal to 1 if IPO firm is listed in a civil law country, and 0 otherwise.
Confidence in Press	Country-specific confidence in press measure from the World Value Survey, orthogonalized to *Shareholder right* and *Earnings opacity*.
Confidence in TV News	Country-specific confidence in TV news measure from the World Value Survey, orthogonalized to *Shareholder right* and *Earnings opacity*.
Earnings opacity	Country-specific earnings earning opacity score, based on Boulton et al. (2011).
Emerging	Dummy variable equal to 1 if IPO firm is listed in an emerging market, and 0 otherwise.
Firm size	Log transformation of total assets of IPO firm (in million US$) at time of listing.
First-day return	IPO first-day closing price minus offer price, scaled by offer price.
Float	Regular shares issued to the public for trading divided by total number of outstanding shares.
GDP per capita growth	Country-specific GDP per capita growth in year of IPO firm listing.
Hot issue market	Average initial return for IPOs issued during the three months prior to month of firm IPO.
Internet censorship	Country-specific internet freedom measure from Freedom House, orthogonalized to *Shareholder right* and *Earnings opacity*.
Internet users	Country-specific proportion of people subscribing to internet in year of IPO firm listing.
IPO size	Total IPO proceeds divided by Total Assets of the IPO firm at time of listing.
Leverage	Total Debt divided by total assets of the IPO firm at time of listing.
Market-to-book	Market-to-Book value of the IPO firm at time of listing.
Market size	Country-specific total value of stock traded divided by GDP at year of IPO listing.
Market turnover	Country-specific turnover ratio of the year of IPO firm listing.
Media coverage	Log transformation of number of times IPO firm is cited in media up to 30 days prior to listing.
Media coverage_Earnings news	Log transformation of number of times news cited in media up to 30 days prior to listing is earnings related.

Media coverage_IPO news	Log transformation of number of times news cited in media up to 30 days prior to listing is IPO related.
Media coverage_Other news	Log transformation of number of times news cited in media up to 30 days prior to listing is something other than IPO/earnings related.
Media coverage_Full article	Log transformation of number of times there is a full article about the IPO firm in media up to 30 days prior to listing.
Media coverage_Other article	Log transformation of number of times there is an article other than a full article *i.e.* newsflash, hot newsflash, press release and tabular material about the IPO firm in the media up to 30 days prior to listing.
Media coverage_Positive news	Log transformation of number of times news cited in media up to 30-days prior to listing has positive sentiment.
Media coverage_Negative news	Log transformation of number of times news cited in media up to 30-days prior to listing has negative sentiment.
Media coverage_Neutral news	Log transformation of number of times news cited in media up to 30-days prior to listing has neutral sentiment.
Media coverage_[-10,-1]	Log transformation of number of times IPO firm is cited in media during days [-10,-1] prior to listing.
Media coverage_[-20,-11]	Log transformation of number of times IPO firm is cited in media during days [-20,-11] prior to listing.
Media coverage_[-30,-21]	Log transformation of number of times IPO firm is cited in media during days [-30,-21] prior to listing.
Media variation	Standard deviation of the transformed event sentiment score among all the news articles about the IPO firm up to 30-days prior to listing.
Newspaper users	Country-specific proportion of people subscribing to newspapers.
Press censorship	Country-specific press freedom measure from Freedom House orthogonalized to *Shareholder right* and *Earnings opacity*.
Profitability	EBIT divided by total assets of the IPO firm at the time of listing.
Proximity to DJ branch	Categorical variable equal to 2 if the IPO firm is headquartered in a metropolitan area with at least one Dow Jones office, equal to 1 if the IPO firm is headquartered in a country with at least one Dow Jones office but in a metropolitan area with no Dow Jones office, and 0 otherwise.
Revision_Vol_AbsRes	Absolute value of the price revision regression residual.
Revision_Vol_SqRes	Square of the price revision regression residual.
ROA	Net income divided by total assets of IPO firm, measured at the end of the first fiscal years after IPO listing.
Security law	Country-specific Securities Law variable for the year of IPO firm listing, based on LaPorta et al. (2006).
Shareholder right	Country-specific Shareholder Rights Index based on Djankov et al. (2008) and Spamann (2010).
Underwriter	Dummy variable equal to 1 if the investment bank underwriting the IPO is in top quartile, and 0 otherwise, based on Lin et al. (2013).
VC back	Dummy variable equal to 1 if the IPO firm is backed by venture capital, and 0 otherwise.

FIGURE 1

Media Coverage and IPO First-Day Return: Univariate Analysis

Figure 1 plots the average IPO first-day return for deciles by media coverage. We divide the sample into deciles of number of times an IPO firm is cited up to 30 days prior to the listing date and plot the mean first-day return for each decile. Variable definitions are presented in Appendix A.

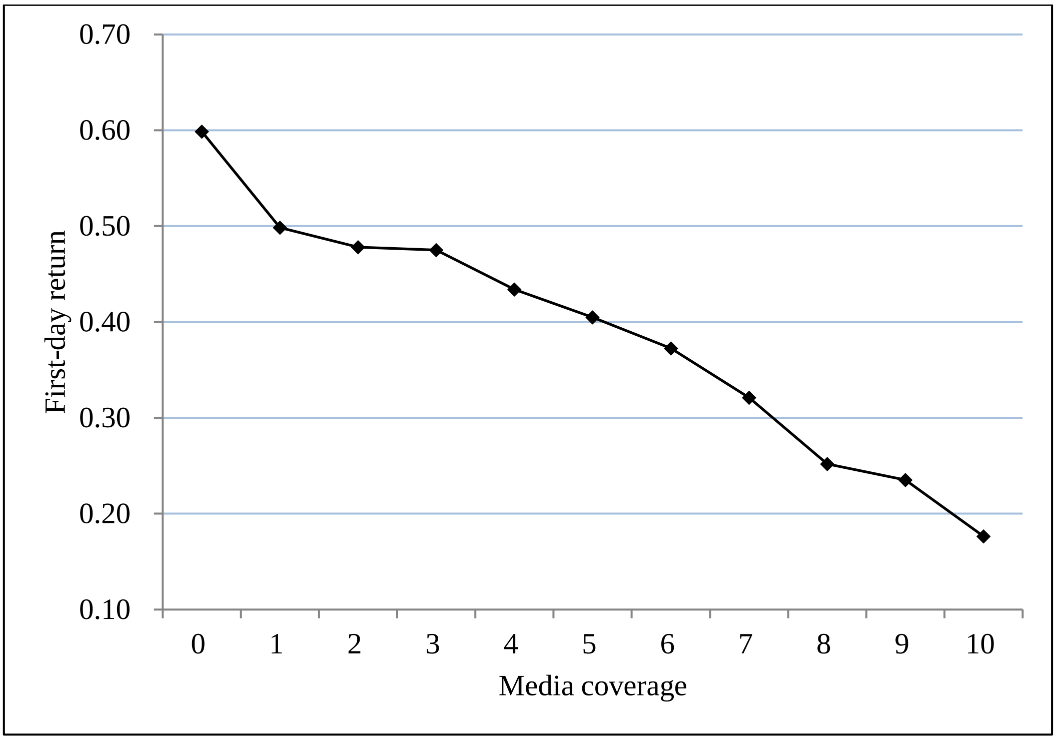

FIGURE 2

Media Coverage and IPO First-Day Return: Semiparametric Regression Analysis

Figure 2 plots the relation between media coverage and IPO first-day returns estimated using Robinson (1988) semi-parametric regression technique. The link function between IPO first-day return and media coverage (solid line) was estimated non-parametrically using kernel method. The explanatory variable was trimmed at 0.1% at the upper tail to mitigate the potential impact of outliers on the non-parametric estimate of the regression curve. The grey area denotes pointwise 95% confidence interval.

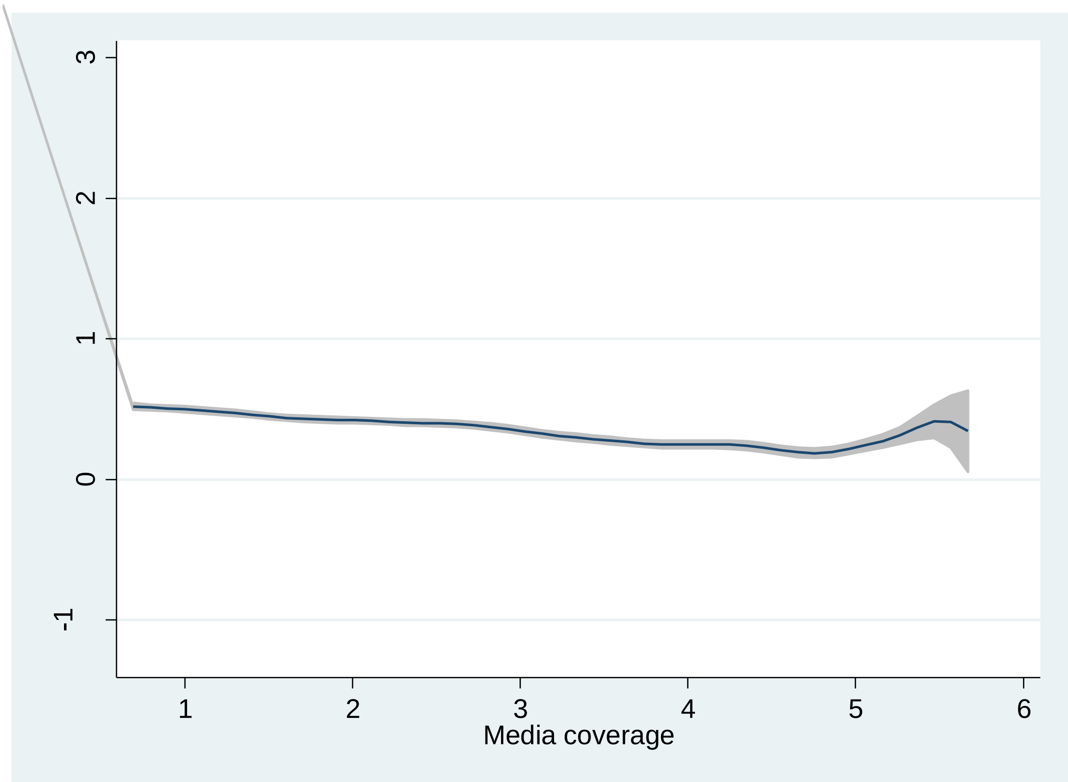

TABLE 1
Sample Distribution and Summary Statistics

Table 1 presents the sample distribution and summary statistics for the variables used in this study. Our sample consists of 11,716 IPOs across 39 countries spanning the period 2000 to 2014. Variable definitions are presented in Appendix A.

Panel A: Country Distribution

Country	No. of IPO	Average First-day Return	Average No. of News Articles
Argentina	10	0.321	12.700
Australia	868	0.322	11.561
Austria	28	0.186	14.714
Belgium	47	0.181	8.936
Brazil	111	0.202	7.919
Canada	578	0.298	13.054
Chile	20	0.240	12.250
China	1,463	0.704	8.648
Denmark	36	0.339	11.250
Finland	17	0.031	17.529
France	314	0.161	13.908
Germany	246	0.111	13.841
Greece	106	0.098	19.406
Hong Kong	452	0.375	10.907
India	460	0.296	12.104
Indonesia	165	0.431	11.739
Ireland	22	0.788	6.409
Italy	131	0.166	9.122
Japan	1,191	0.470	13.219
Malaysia	449	0.283	16.744
Mexico	17	0.323	8.529
Netherland	52	0.328	19.673
New Zealand	38	0.130	13.079
Norway	72	0.283	12.417
Philippines	30	0.425	12.433
Poland	150	0.486	11.687
Portugal	10	-0.089	17.900
Russia	49	0.148	10.673
Singapore	364	0.320	12.665
Spain	39	0.158	12.487
Sweden	70	0.187	15.457
Switzerland	53	0.129	9.962
Taiwan	739	0.186	24.553
Thailand	256	0.252	12.395
Turkey	46	0.025	19.065
South Africa	23	0.487	17.087
South Korea	601	0.407	13.388
U.K.	934	0.397	6.687
U.S.	1,459	0.373	22.600
Total	11,716	0.387	14.058

Panel B: Year Distribution

Year	No. of IPOs	Average First-day Return	Average No. of News Articles
2000	1,118	0.204	10.240
2001	625	0.256	12.894
2002	598	0.260	14.237
2003	627	0.344	11.764
2004	1,147	0.343	13.428
2005	1,134	0.310	12.326
2006	1,303	0.317	12.450
2007	1,510	0.366	14.350
2008	566	0.272	11.733
2009	365	0.624	21.858
2010	750	0.720	15.231
2011	652	0.652	16.268
2012	453	0.542	19.119
2013	475	0.592	17.091
2014	393	0.616	19.802
Total	11,716	0.387	14.058

Panel C: Summary Statistics

	Mean	Std. Dev.	5%	Median	95%
First-day return	0.387	0.560	-0.200	0.221	1.600
Media coverage	1.755	1.373	0.000	1.609	4.304
Firm size	4.811	1.779	2.046	4.705	7.874
Profitability	0.042	0.170	-0.380	0.054	0.293
Leverage	0.121	0.191	0.006	0.033	0.542
Market-to-book	3.682	4.545	0.694	2.380	10.730
Asset turnover	0.769	0.805	0.000	0.553	2.426
Bookbuilding	0.588	0.492	0.000	1.000	1.000
GDP per capita growth	0.039	0.035	-0.002	0.026	0.099
Market size	1.195	0.941	0.229	1.022	3.052
Market turnover	1.164	0.602	0.343	1.040	2.165

TABLE 2

Media Coverage and IPO First-day Return: Baseline Regression Results

Table 2 presents the regression results for the relation between media coverage and IPO first-day return. Our sample consists of 11,716 IPOs across 39 countries spanning the period 2000 to 2014. The regressions are performed by OLS, with *t*-statistics computed using standard errors robust to heteroskedasticity. Constant, industry fixed effects based on Kenneth French's 10-industry classification, year of listing fixed effects, and country of listing fixed effects are included in all the regressions. Variable definitions are presented in Appendix A.

Dependent Variable:	*First-day return* (1)		*First-day return* (2)		*First-day return* (3)	
	Co-eff.	*t*-stat.	Co-eff.	*t*-stat.	Co-eff.	*t*-stat.
Media coverage	-0.040	-11.68	-0.038	-11.24	-0.037	-11.04
Firm size			0.006	1.70	0.005	1.74
Profitability			0.078	2.37	0.080	2.43
Leverage			0.258	7.68	0.263	7.79
Market-to-book			0.003	2.64	0.003	2.83
Asset turnover			-0.041	-5.54	-0.041	-5.56
Bookbuilding			-0.142	-10.82	-0.149	-11.32
GDP per capita growth					0.035	0.10
Market size					0.058	4.49
Market turnover					0.018	0.93
Industry FE	Yes		Yes		Yes	
Year FE	Yes		Yes		Yes	
Country FE	Yes		Yes		Yes	
Observations	11,716		11,716		11,716	
Adjusted R^2	0.174		0.190		0.194	

TABLE 3

Media Coverage and IPO First-day Return: Country and Year Regressions

Table 3 presents country-by-country and year-by-year regression results for the relation between media coverage and IPO first-day return. For brevity, the table only reports the coefficient of media coverage. Our sample consists of 11,716 IPOs across 39 countries spanning the period 2000 to 2014. In Panel A, we only include the regression results for countries for which we have at least 50 or more IPOs over the sample period. The regressions are performed by OLS, with *t*-statistics computed using standard errors robust to heteroskedasticity. Control variables and fixed effects are included in all the regressions but not tabulated for brevity. Variable definitions are presented in Appendix A.

Panel A: Country Regressions

Country	Co-eff.	*t*-stat.	Adjusted R^2	Observations
Australia	-0.027	-2.30	0.251	868
Brazil	-0.066	-1.87	0.366	111
Canada	-0.050	-2.62	0.140	578
China	-0.128	-10.01	0.334	1,463
France	-0.053	-3.67	0.244	314
Germany	-0.028	-0.98	0.299	246
Greece	0.000	0.01	0.413	106
Hong Kong	-0.015	-0.74	0.293	452
Indonesia	-0.091	-2.15	0.327	165
India	-0.046	-2.67	0.432	460
Italy	0.006	0.18	0.436	131
Japan	-0.061	-6.12	0.215	1,191
Malaysia	-0.071	-5.51	0.257	449
Netherland	-0.127	-1.81	0.494	52
Norway	0.008	0.22	0.680	72
Poland	-0.114	-3.18	0.473	150
Singapore	-0.044	-2.69	0.277	364
South Korea	-0.069	-4.66	0.230	601
Sweden	-0.067	-1.79	0.353	70
Switzerland	0.108	1.49	0.337	53
Thailand	-0.026	-1.69	0.268	256
Taiwan	-0.011	-1.30	0.213	739
U.K.	-0.026	-1.99	0.282	934
U.S.	-0.061	-4.14	0.226	1,459

Panel B: Year Regressions

Year	Co-eff.	*t*-stat.	Adjusted R^2	Observations
2000	-0.017	-1.46	0.337	1,118
2001	-0.028	-2.17	0.269	625
2002	-0.015	-1.27	0.210	598
2003	-0.029	-2.21	0.225	627
2004	-0.039	-3.85	0.223	1,147
2005	-0.021	-2.19	0.284	1,134
2006	-0.032	-3.31	0.169	1,303
2007	-0.018	-1.89	0.252	1,510
2008	0.005	0.34	0.229	566
2009	-0.114	-4.34	0.359	365
2010	-0.131	-7.87	0.427	750
2011	-0.081	-4.50	0.357	652
2012	-0.052	-2.27	0.326	453
2013	-0.053	-1.98	0.299	475
2014	-0.041	-1.73	0.322	393

TABLE 4

Media Coverage and IPO First-Day Return: Robustness Checks

Table 4 presents the regression results for various robustness checks for the relation between media coverage and IPO first-day return. For brevity, the table only reports the coefficient of media coverage. Our sample consists of up to 11,716 IPOs across 39 countries depending upon the model specification spanning the period 2000 to 2014. The regressions are performed by OLS, with *t*-statistics computed using standard errors robust to heteroskedasticity. Control variables, constant, industry fixed effects based on Kenneth French's 10-industry classification, year of listing fixed effects, and country of listing fixed effects are included in all the regressions but not tabulated for brevity. Variable definitions are presented in Appendix A.

Panel A: Alternative measures of IPO initial return		
(1) IPO return over 1-week after listing		
Dependent variable:	*One-week return*	
	Co-eff.	*t*-stat.
Media coverage	-0.048	-4.46
(2) IPO return over 2 weeks after listing		
Dependent variable:	*Two-week return*	
	Co-eff.	*t*-stat.
Media coverage	-0.047	-3.44
Panel B: Alternative measures of media coverage		
(1) News articles in 60 days prior to IPO		
Dependent variable:	*First-day return*	
	Co-eff.	*t*-stat.
Media coverage	-0.036	-10.64
(2) News articles in 90 days prior to IPO		
Dependent variable:	*First-day return*	
	Co-eff.	*t*-stat.
Media coverage	-0.033	-10.15
(3) Media data from Factiva		
Dependent variable:	*First-day return*	
	Co-eff.	*t*-stat.
Media coverage	-0.044	-7.87
Panel C: Additional control variables		
(1) Include additional control variables – IPO float, age, size, cash balance, advertising intensity and hot issue market		
Dependent variable:	*First-day return*	
	Co-eff.	*t*-stat.
Media coverage	-0.045	-11.08
Panel D: Alternative sample specifications		
(1) Exclude IPOs with zero media coverage		
Dependent variable:	*First-day return*	
	Co-eff.	*t*-stat.
Media coverage	-0.074	-19.18
Panel E: Decomposition of IPO initial return		
(1) Primary market return		
Dependent variable:	*Primary market return*	
	Co-eff.	*t*-stat.
Media coverage	-0.046	-5.91
(4) Secondary market return		
Dependent variable:	*Secondary market return*	
	Co-eff.	*t*-stat.
Media coverage	-0.002	-0.68

TABLE 5

Media Coverage and IPO First-Day Return: National Media Strikes as Exogenous Shocks

Table 5 presents the regression results using national media strikes as exogenous shocks to media coverage. Our sample consists of 581 IPOs that were listed between 2000 and 2014 in the countries that experienced at least one media strike during this period. *Strike* is a dummy variable equal to one if the IPO has at least one media strike taking place in the same country during the 30-day window prior to the IPO date, and zero otherwise. The regressions are performed by OLS, with *t*-statistics computed using standard errors robust to heteroskedasticity. Constant, industry fixed effects based on Kenneth French's 10-industry classification, year of listing fixed effects, and country of listing fixed effects are included. Variable definitions are presented in Appendix A.

Panel A: Comparison of IPOs with and without media strike

	Strike = 0 (n = 519)	*Strike* = 1 (n = 62)	Difference	*t*-stat.
Mean number of news articles	14.202	9.726	4.476	2.09

Panel B: Regression analysis on the effect of media strikes

Dependent Variable:	*First-day return* (1)	
	Co-eff.	*t*-stat.
Strike	0.160	2.18
Firm size	-0.047	-3.32
Profitability	0.178	1.47
Leverage	0.569	3.10
Market-to-book	-0.003	-0.74
Asset turnover	-0.046	-1.98
Bookbuilding	0.037	0.76
GDP per capita growth	6.148	1.79
Market size	-0.507	-2.49
Market turnover	0.228	1.70
Industry FE	Yes	
Year FE	Yes	
Country FE	Yes	
Observations	581	
Adjusted R^2	0.199	

TABLE 6

Media Coverage and IPO First-Day Return: Instrumental Variable Estimation

Table 6 presents the results for the instrumental variable estimation, with *t*-statistics computed using standard errors robust to heteroskedasticity. Our sample consists of 11,716 IPOs across 39 countries spanning the period 2000 to 2014. Constant, industry fixed effects based on Kenneth French's 10-industry classification, year of listing fixed effects, and country of listing fixed effects are included in all the regressions. Variable definitions are presented in Appendix A.

Dependent Variable:	*Media coverage* (1)		*First-day return* (2)	
	Co-eff.	*t*-stat.	Co-eff.	*t*-stat.
Proximity to DJ branch	0.054	2.62		
Media coverage			-0.375	-3.35
Firm size	0.031	3.68	-0.014	-1.86
Profitability	-0.140	-1.76	0.116	1.80
Leverage	-0.093	-1.44	0.241	2.97
Market-to-book	-0.002	-0.63	0.003	1.57
Asset turnover	0.023	1.29	-0.035	-1.81
Bookbuilding	0.117	3.60	-0.121	-2.32
GDP per capita growth	2.280	2.57	0.579	0.32
Market size	-0.124	-4.04	0.028	0.29
Market turnover	0.088	2.08	0.039	0.54
Industry FE	Yes		Yes	
Year FE	Yes		Yes	
Country FE	Yes		Yes	
Observations	11,716		11,716	
Adjusted R^2	0.190		-	

TABLE 7

Media Coverage and IPO First-Day Return: The Moderating Effect of Country-Level Institutions

Table 7 presents the regression results for the effects of country-level institutions on the relation between media coverage and IPO first-day return. Our sample consists of up to 11,716 IPOs across 39 countries depending upon the model specification spanning the period 2000 to 2014. The regressions are performed by OLS, with *t*-statistics computed using standard errors robust to heteroskedasticity. Constant, industry fixed effects based on Kenneth French's 10-industry classification and year of listing fixed effects are included in all the regressions. Variable definitions are presented in Appendix A.

Dependent Variable:	*First-day return* (1)		*First-day return* (2)		*First-day return* (3)		*First-day return* (4)		*First-day return* (5)		*First-day return* (6)	
	Co-eff.	*t*-stat.	Co-eff.	*t*-stat.	Co-eff.	*t*-stat.	Co-eff.	*t*-stat.	Co-eff.	*t*-stat.	Co-eff.	*t*-stat.
Media coverage	-0.128	-7.86	-0.058	-7.97	-0.122	-9.59	-0.078	-3.99	-0.009	-1.88	-0.031	-7.82
*Media coverage*Earning opacity*	-0.286	-10.10										
Earnings opacity	0.247	4.68										
*Media coverage*Accounting Conservatism*			0.148	5.47								
Accounting Conservatism			-0.551	-7.41								
*Media coverage*Shareholder rights*					0.021	6.55						
Shareholder right					-0.099	-2.30						
*Media coverage*Security law*							0.439	3.64				
Security law							0.209	0.76				
*Media coverage*Civil law*									-0.060	-8.68		
Civil law									0.151	7.60		
*Media coverage*Emerging*											-0.040	-4.62
Emerging											0.120	5.15
Firm size	0.021	6.15	-0.002	-0.68	0.009	2.56	0.019	5.87	0.016	4.77	0.014	4.30
Profitability	-0.012	-0.37	0.053	1.60	0.093	2.82	0.076	2.35	0.069	2.10	0.045	1.38
Leverage	0.306	8.82	0.345	9.69	0.285	8.40	0.284	8.40	0.291	8.57	0.278	8.15
Market-to-book	0.004	3.40	0.005	4.05	0.004	3.60	0.004	3.47	0.004	3.58	0.004	3.87
Asset turnover	-0.035	-4.70	-0.021	-2.76	-0.031	-4.08	-0.036	-4.71	-0.038	-5.07	-0.031	-4.09
Bookbuilding	-0.089	-7.48	-0.029	-2.42	-0.067	-5.81	-0.058	-4.93	-0.071	-6.04	-0.059	-5.08
GDP per capita growth	-0.596	-2.70	-1.914	-9.49	0.072	0.39	1.444	8.35	1.250	7.28	0.245	1.36
Market size	0.035	4.74	0.055	6.97	0.044	5.93	0.013	1.57	0.036	4.34	0.042	5.75
Market turnover	0.080	6.82	0.006	0.52	0.039	3.27	0.083	7.12	0.058	4.79	0.090	7.71
Industry FE	Yes		Yes		Yes		Yes		Yes		Yes	
Year FE	Yes		Yes		Yes		Yes		Yes		Yes	
Country FE	No		No		No		No		No		No	
Observations	11,497		10,054		11,716		11,716		11,716		11,716	
Adjusted R^2	0.141		0.112		0.134		0.119		0.124		0.130	

TABLE 8

Media Coverage and IPO First-Day Return: The Moderating Effect of Country-Level Media Characteristics

Table 8 presents the regression results for the effects of country-level media characteristics on the relation between media coverage and IPO first-day return. Our sample consists of up to 11,716 IPOs across 39 countries depending upon the model specification spanning the period 2000 to 2014. The regressions are performed by OLS, with *t*-statistics computed using standard errors robust to heteroskedasticity. Constant, industry fixed effects based on Kenneth French's 10-industry classification and year of listing fixed effects are included in all the regressions. Variable definitions are presented in Appendix A.

Dependent Variable:	*First-day return* (1)		*First-day return* (2)		*First-day return* (3)		*First-day return* (4)		*First-day return* (5)		*First-day return* (6)	
	Co-eff.	*t*-stat.	Co-eff.	*t*-stat.	Co-eff.	*t*-stat.	Co-eff.	*t*-stat.	Co-eff.	*t*-stat.	Co-eff.	*t*-stat.
Media coverage	-0.026	-3.60	-0.041	-4.59	-0.040	-11.42	-0.045	-10.79	-0.038	-10.84	-0.036	-10.26
*Media coverage*Newspaper users*	-0.066	-2.80										
Newspaper users	0.042	0.67										
*Media coverage*Internet users*			-0.024	-2.15								
Internet users			0.043	0.90								
*Media coverage*Press censorship*					0.173	16.85						
Press censorship					-0.032	-7.75						
*Media coverage*Internet censorship*							0.215	13.59				
Internet censorship							-0.043	-8.55				
*Media coverage*Confidence in press*									-0.034	-9.44		
Confidence in press									0.149	15.91		
*Media coverage*Confidence in TV News*											-0.039	-10.86
Confidence in TV News											0.168	17.60
Firm size	0.018	5.47	0.012	3.62	0.014	4.34	0.011	3.05	0.022	6.51	0.023	6.85
Profitability	0.076	2.35	0.062	1.85	0.011	0.32	0.082	2.23	-0.027	-0.82	-0.023	-0.69
Leverage	0.280	8.27	0.276	8.01	0.295	8.57	0.251	6.86	0.281	7.99	0.276	7.90
Market-to-book	0.004	3.46	0.004	3.55	0.005	3.96	0.005	3.78	0.005	3.92	0.005	3.97
Asset turnover	-0.035	-4.59	-0.037	-4.68	-0.023	-3.03	-0.025	-2.91	-0.035	-4.58	-0.034	-4.48
Bookbuilding	-0.059	-5.10	-0.072	-5.99	-0.046	-3.99	-0.060	-4.57	-0.094	-7.84	-0.092	-7.69
GDP per capita growth	1.243	7.26	2.477	10.28	-0.693	-3.70	-0.735	-1.99	0.782	4.63	0.544	3.20
Market size	0.022	3.01	0.010	1.36	0.030	4.23	0.090	6.18	0.044	6.03	0.035	4.84
Market turnover	0.080	6.93	0.104	8.73	0.092	7.93	0.015	0.88	0.074	6.32	0.081	6.85
Industry FE	Yes		Yes		Yes		Yes		Yes		Yes	
Year FE	Yes		Yes		Yes		Yes		Yes		Yes	
Country FE	No		No		No		No		No		No	
Observations	11,716		10,977		11,497		9,352		11,248		11,248	
Adjusted R^2	0.119		0.139		0.149		0.163		0.146		0.150	

TABLE 9
Media Coverage and IPO First-Day Return: The Moderating Effect of IPO Certification

Table 9 presents the regression results for the effect of IPO certification on the relation between media coverage and IPO first-day return. Our sample consists of 10,681 IPOs for which we have data on certification characteristics across 39 countries spanning the period 2000 to 2014. The regressions are performed by OLS, with *t*-statistics computed using standard errors robust to heteroskedasticity. Constant, industry fixed effects based on Kenneth French's 10-industry classification, year of listing fixed effects, and country of listing fixed effects are included in all the regressions. Variable definitions are presented in Appendix A.

Dependent Variable:	*First-day return* (1)		*First-day return* (2)		*First-day return* (3)	
	Co-eff.	*t*-stat.	Co-eff.	*t*-stat.	Co-eff.	*t*-stat.
Media coverage	-0.053	-12.82	-0.053	-12.08	-0.049	-11.75
*Media coverage*VC back*	0.041	4.92				
VC back	-0.060	-2.87				
*Media coverage*Big 4 auditor*			0.093	3.87		
Big 4 auditor			-0.072	-3.66		
*Media coverage*Underwriter*					0.052	2.54
Underwriter					-0.024	-1.19
Firm size	0.004	1.00	0.004	0.98	0.003	0.84
Profitability	0.087	2.50	0.075	2.15	0.075	2.15
Leverage	0.267	7.81	0.261	7.61	0.258	7.50
Market-to-book	0.003	2.69	0.003	2.64	0.003	2.74
Asset turnover	-0.047	-6.16	-0.049	-6.35	-0.048	-6.23
Bookbuilding	-0.153	-11.06	-0.151	-10.90	-0.153	-11.07
GDP per capita growth	-0.142	-0.38	-0.118	-0.32	-0.169	-0.45
Market size	0.060	4.56	0.060	4.49	0.061	4.61
Market turnover	0.010	0.48	0.010	0.47	0.008	0.41
Industry FE	Yes		Yes		Yes	
Year FE	Yes		Yes		Yes	
Country FE	Yes		Yes		Yes	
Observations	10,681		10,681		10,681	
Adjusted R^2	0.196		0.195		0.195	

TABLE 10
Media Coverage and IPO Information Asymmetry

Table 10 presents the results of the test of information asymmetry channel. Our sample consists of 4,794 IPOs that are priced using the bookbuilding approach and for which we have price revision data available across 39 countries spanning the period 2000 to 2014. The regressions are performed by OLS, with *t*-statistics computed using standard errors robust to heteroskedasticity. Constant, industry fixed effects based on Kenneth French's 10-industry classification, year of listing fixed effects, and country of listing fixed effects are included in all the regressions. Variable definitions are presented in Appendix A.

Dependent Variable:	*First-day return*		*First-day return*		*Revision_Vol_ SqRes*		*Revision_Vol_ AbsRes*	
	(1)		(2)		(3)		(4)	
	Co-eff.	*t*-stat.	Co-eff.	*t*-stat.	Co-eff.	*t*-stat.	Co-eff.	*t*-stat.
Revision_Vol_SqRes	1.264	2.66						
Revision_Vol_AbsRes			0.361	2.90				
Media coverage					-0.001	-3.30	-0.003	-2.83
Firm size	-0.006	-1.04	-0.006	-1.00	-0.001	-2.70	-0.003	-3.53
Profitability	0.103	2.01	0.103	2.03	0.003	1.24	0.006	0.92
Leverage	0.187	4.35	0.183	4.27	0.002	1.61	0.018	3.40
Market-to-book	0.005	2.74	0.005	2.77	0.000	0.88	0.000	0.22
Asset turnover	-0.082	-7.97	-0.082	-7.96	-0.001	-2.42	-0.004	-2.37
GDP per capita growth	0.312	0.51	0.313	0.51	-0.026	-1.11	-0.093	-1.12
Market size	0.049	2.01	0.049	2.02	-0.003	-2.52	-0.010	-2.86
Market turnover	0.069	1.93	0.069	1.94	0.002	1.29	0.005	0.94
Industry FE	Yes		Yes		Yes		Yes	
Year FE	Yes		Yes		Yes		Yes	
Country FE	Yes		Yes		Yes		Yes	
Observations	4,794		4,794		4,794		4,794	
Adjusted R^2	0.187		0.187		0.073		0.077	

TABLE 11
Media Coverage and IPO First-Day Return: News Characteristics

Table 11 presents the regression results for the effect of news characteristics on the relation between media coverage and IPO first-day return. Our sample consists of 11,716 IPOs across 39 countries spanning the period 2000 to 2014. The regressions are performed by OLS, with *t*-statistics computed using standard errors robust to heteroskedasticity. Constant, industry fixed effects, year fixed effects, and country fixed effects, are included in all the regressions. Variable definitions are presented in Appendix A.

Dependent Variable:	*First-day return* (1)		*First-day return* (2)		*First-day return* (3)		*First-day return* (4)		*First-day return* (5)	
	Co-eff.	*t*-stat.	Co-eff.	*t*-stat.	Co-eff.	*t*-stat.	Co-eff.	*t*-stat.	Co-eff.	*t*-stat.
Media coverage_Earnings news	-0.033	-6.65								
Media coverage_IPO news	-0.022	-3.90								
Media coverage_Other news	-0.017	-3.51								
Media coverage_Full article			-0.031	-6.17						
Media coverage_Other article			-0.016	-3.36						
Media coverage					-0.081	-9.08				
*Media coverage*Media variation*					0.109	2.64				
Media variation					0.163	2.02				
Media coverage_Positive news							-0.040	-6.67		
Media coverage_Negative news							-0.038	-9.41		
Media coverage_Neutral news							-0.010	-1.85		
Media coverage_[-10,-1]									-0.025	-3.07
Media coverage_[-20,-11]									-0.015	-2.58
Media coverage_[-30,-21]									-0.009	-1.70
Firm size	0.004	1.16	0.005	1.28	-0.003	-0.84	-0.003	-0.93	0.005	1.31
Profitability	0.078	2.39	0.080	2.43	0.072	2.20	0.067	2.04	0.086	2.61
Leverage	0.263	7.79	0.263	7.80	0.229	6.86	0.234	7.00	0.264	7.80
Market-to-book	0.003	2.91	0.003	2.88	0.004	3.58	0.004	3.61	0.003	2.93
Asset turnover	-0.040	-5.48	-0.040	-5.52	-0.048	-6.50	-0.047	-6.36	-0.041	-5.56
Bookbuilding	-0.149	-11.29	-0.150	-11.38	-0.152	-11.36	-0.151	-11.29	-0.152	-11.49
GDP per capita growth	0.032	0.09	0.020	0.06	-0.375	-1.06	-0.401	-1.14	-0.016	-0.05
Market size	0.057	4.46	0.059	4.56	0.041	3.30	0.042	3.39	0.058	4.49
Market turnover	0.018	0.94	0.017	0.91	0.009	0.49	0.008	0.40	0.020	1.02

TABLE 11 (continued)

Industry FE	Yes	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes	Yes
Country FE	Yes	Yes	Yes	Yes	Yes
Observations	11,716	11,716	9,431	9,431	11,716
Adjusted R^2	0.197	0.195	0.188	0.189	0.191

TABLE 12

Media Coverage and IPO First-day Return: Reconciliation with U.S. Evidence

Table 12 presents the regression results of the relation between media coverage and IPO first-day return in the U.S. market using Factiva data. Our sample consists of 2,348 IPOs in the U.S. spanning the period 1993 to 2014. The regressions are performed by OLS, with *t*-statistics computed using standard errors robust to heteroskedasticity. Constant, industry fixed effects based on Kenneth French's 10-industry classification, year of listing fixed effects, and country of listing fixed effects are included in all the regressions. Variable definitions are presented in Appendix A.

Panel A. Main results

	Sample Period 1993-2000		Sample Period 2001-2014	
Dependent Variable:	*First-day return* (1)		*First-day return* (2)	
	Co-eff.	*t*-stat.	Co-eff.	*t*-stat.
Media coverage	0.110	4.35	-0.063	-5.13
Firm size	-0.020	-1.81	0.028	2.22
Profitability	0.317	5.67	0.013	0.17
Leverage	-0.028	-0.35	0.079	1.26
Market-to-book	0.034	2.48	0.009	2.97
Asset turnover	-0.110	-3.66	-0.101	-5.22
Bookbuilding	-1.089	-2.26	0.203	3.06
Industry FE	Yes		Yes	
Year FE	Yes		Yes	
Observations	1,289		1,059	
Adjusted R^2	0.255		0.132	

Panel B. Robustness tests

	Sample Period 1993-2000		Sample Period 2001-2014	
Dependent Variable:	*First-day return* (1)		*First-day return* (2)	
(1) Exclude IPOs in the IT industry				
	Co-eff.	*t*-stat.	Co-eff.	*t*-stat.
Media coverage	0.112	4.23	-0.062	-4.89
(2) News articles in 60 days prior to IPO				
	Co-eff.	*t*-stat.	Co-eff.	*t*-stat.
Media coverage	0.139	4.38	-0.068	-5.26
(3) News articles in 90 days prior to IPO				
	Co-eff.	*t*-stat.	Co-eff.	*t*-stat.
Media coverage	0.153	4.40	-0.082	-5.89
(4) Different types of articles				
	Co-eff.	*t*-stat.	Co-eff.	*t*-stat.
Media coverage_Full article	0.231	2.53	-0.089	-4.43
Media coverage_Other article	0.017	0.19	-0.036	-2.05